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# The euro impact on trade. Long run evidence with structural breaks<sup>\*</sup>

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## Abstract

In this paper we present new evidence on the euro effect on trade. We use a data set containing all bilateral combinations in a panel of 26 OECD countries during the period 1967-2008. From a methodological point of view, we implement a new generation of tests that allow solving some of the problems derived from the non-stationary nature of the data. To this aim we apply panel tests that account for the presence of cross-section dependence as well as discontinuities in the non-stationary panel data. We test for cointegration between the variables using panel cointegration tests, especially the ones proposed by Banerjee and Carrión-i-Silvestre (2010). We also efficiently estimate the long-run relationships using the CUP-BC and CUP-FM estimators proposed in Bai et al. (2009). We argue that, after controlling for cross-section dependence and deterministic trends and breaks in trade integration, the euro appears to generate lower trade effects than predicted in previous studies.

**JEL classification numbers:** C12, C22, F15, F10.

**Key words:** Gravity models; trade; panel cointegration; common factors; structural breaks, cross-section dependence.

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## **I. Introduction**

The introduction of the euro has raised a new interest in measuring the impact of currency unions (CU) on trade flows. The very high estimates of trade induced by the creation of monetary unions found in the seminal papers by Rose (2000) and Frankel and Rose (2002) has led to the concept of ‘endogeneity’ of Optimum Currency Areas (OCA) that means for the euro area that, even if the European Monetary Union (EMU) was not created as an OCA, it could be progressing in that direction (Frankel and Rose, 1998). Recent research surveyed by Rose and Stanley (2005) and Rose (2008) suggests that the introduction of the euro still has a sizable and statistically significant effect on trade among EMU members. Taking together all these estimates imply that EMU has increased trade by about 8%-23% percent in its first years of existence. This issue can be very relevant for prospective new members of EMU.

In 1999 eleven countries of the EU adopted the euro as a common currency while Greece entered in 2001. Since then, also Slovenia, Cyprus, Malta, Slovakia and Estonia have joined the euro area while other members of the EU are ‘waiting and seeing’, the so-called derogation countries. Moreover, the introduction of the euro was preceded by other stages of economic integration (Customs Union, European Monetary System and the Single Market), so the EMU effect has to be analyzed as an on-going process with a time dimension. It might be interesting to investigate whether there is an additional benefit of a common currency over (relative) exchange rate stability. As pointed out by Faruquee (2004) the central questions at stake are the following: first, to ascertain the effects of EMU on the area’s trade flows; second, to analyze the evolution of the trade effects through time, and finally, to measure the distribution of trade effects among member states.

In this paper we have tried to overcome some of the main flaws found in the standard

empirical literature and recently outlined by Eicher and Henn (2011). First, Baldwin and Taglioni's (2006, BT henceforth) critiques regarding the proper specification of gravity models in large panels to prevent omitted variable bias point out the need to simultaneously account for multilateral resistance and unobserved bilateral heterogeneity. We have accounted for BT's critiques in the specification of the model as well as in the definition of the variables included in the estimation of the gravity model.

Second, more recently, Fidrmuc (2009) and Bun and Klaasen (2007) have outlined the importance of considering the possible non stationary nature of the variables included in the gravity equation, as well as the cross-sectional correlation between the elements (countries) of the panel, both aspects normally neglected in the empirical applications. While initially the literature overlooked some crucial econometric issues regarding non-stationary series in panel estimation, more recent works have taken into account these aspects using non-stationary panel data techniques. A sizeable literature has been developing along these lines, but none of these works explicitly deals with the issue of cross-section dependence with the exception of Gengenbach (2009).

Third, Bun and Klaasen (2007) have stated that models measuring the effect of the euro on trade have omitted some variables, causing an upward bias in the trade benefits earlier estimated. They find that the longer the data period considered, the higher the euro effect estimate. Thus this might be due to some misspecification of the time-series characteristics of the variables involved, namely the trends in trade flows over time. To correct for this bias they add a time-trend to their specification and allow it to have heterogeneous coefficients across country-pairs. Then they estimate long run relationships using first-generation panel cointegration techniques, that is, without considering dependence in the cross-section dimension.

Therefore, in this paper we try to provide new evidence on the effect of the euro using a data set that contains information on all bilateral combinations in a panel of 26 countries

covering the period 1967-2008. We implement a new generation of tests that allows us to solve some of the problems derived from the non-stationary nature of the data used in gravitational equations. More specifically, we implement the panel unit root and stationary tests proposed by Pesaran (2004, 2007) and Bai and Ng (2004) to test for the presence of cross-section dependence as well as discontinuities in the non-stationary series. We then test for cointegration between the variables using panel cointegration tests, with a special emphasis in the one proposed by Banerjee and Carrión-i-Silvestre (2010). Finally, we apply the continuously updated estimator (CUP) of Bai *et al.* (2009) to efficiently estimate the regression coefficients. The results obtained are in line with Bun and Klaassen (2007) confirming a smaller euro effect than in other research papers, like for instance, Gil-Pareja *et al* (2008), where cross-section dependence and the non-stationary nature of the variables are not accounted for.

The paper is organized as follows. Section 2 discusses the empirical literature on CU and trade, emphasizing the econometric approaches based on the gravity model. Section 3 presents a new econometric approach that overcomes some of the problems present in the current literature, describes the data and discusses the empirical results. A final section concludes.

## **II. Previous studies and criticisms to the empirical application of the gravity equation to measuring the euro effect on trade**

The literature examining the impact of CU on trade is a burgeoning field of research. All in all, the diversity of existing estimates indicates the potential bias inherent in applied specifications. Although in the beginning the gravity model was criticized for its lack of theoretical

underpinnings, now rests on a solid theoretical background.<sup>1</sup> Therefore, as stated in Westerlund and Wilhelmsson (2009) the focus of this line of research has shifted from its theoretical soundness towards the estimation techniques used.

The econometric approach has changed over time as a result of a feedback process between theory and empirics. In this abundant literature, the traditional approach has been to use *cross-section data*. However, it is generally accepted that the results obtained were suffering from a bias, as the heterogeneity among countries was not properly controlled for. Thus, Rose's (2000) initial estimates in a cross-sectional study suggested a tripling of trade. This result was quite striking, and as quoted by Faruquee (2004), is at odds with the related literature that typically finds very little negative impact of exchange rate volatility on trade. Not surprisingly, Rose's findings have received substantial revisions, and subsequent analysis generally finds a smaller (albeit still sizable) effect of CU membership on trade. There are different reasons that make the implication of Rose (2000) work unclear. First, the sample countries were mostly smaller and poorer, not including the EMU ones. This has led to question whether the results apply to bigger countries such as the EMU members. Second, the cross-sectional analysis included in Rose (2000) provides a comparative benchmark across members of a monetary union against third countries but the most relevant issue about EMU is the possible change in the level of trade for its member over time, before and after the introduction of the single currency. In order to solve this problem, a second string of literature started to use *panel data estimation techniques*, which permits more general types of heterogeneity.<sup>2</sup> However, BT define what they call in this context 'the gold medal error', also known as the 'Anderson-van Wincoop (A-vW) misinterpretation' in the sense that A-vW

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<sup>1</sup>See, for instance, Feenstra *et al.* (2001).

<sup>2</sup>Moreover, as clearly explained by Westerlund and Wilhelmsson (2009), if we desire to measure the impact of a currency union on exports (which is the relevant case in this paper), while simultaneously controlling for country-pair propensity to trade, it is easier under a panel data framework by means of a country-pair fixed effect term. For a single cross-section, these controls can only depend on observed country-pair attributes such as common language, and estimates can thus be biased if there is additionally an unobserved component to the country-pair propensity to trade.

developed a cross-section estimation technique to control for omitted variables with pair fixed effects.<sup>3</sup> However, this technique has been generalized to the panel data framework by many authors without considering the time dimension (see, for example, Glick and Rose, 2002 or Flam and Nordstrom, 2006). Country dummies (for exporters and importers) only remove the average impact leaving the time dimension in the residuals, which leads to biased results. Therefore, time-invariant country dummies are not enough and a proper treatment of the time dimension is needed. Moreover, BT also stress the importance of an omitted variable bias when the empirical specification does not account for unobserved determinants of bilateral trading relationships. They suggest the inclusion of time varying fixed effects in the specification. However, if doing so, we would not be able to explore cointegration between GDP and exports, since the time varying fixed effects would absorb GDP. Instead of that, and following Bun and Klaasen (2007), we include in our specification a country-pair specific time trend which captures all the unobserved heterogeneity through time, as well as country specific fixed effects. Furthermore, the application of cointegration techniques implies the proper treatment of the time dimension, since it takes into account the long-run relationships among variables.

In addition to the above-mentioned specification caveats, BT pointed out two additional minor problems, coined as ‘silver’ and ‘bronze’ medal errors. The silver medal error concerns the definition of the dependent variable. As BT point out, the gravity equation is an expenditure function that explains the value of spending by one nation on the goods produced by another nation; it explains uni-directional bilateral trade. Most gravity models, however, work with the average of the two way exports and frequently the averaging procedure is wrong. The problem arises when authors use the log of the sum instead of the sum of the logs in the bilateral trade term. The silver medal mistake will create no bias if bilateral trade is balanced. However, if nations in a currency union tend to have larger than usual bi-

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<sup>3</sup> See Anderson and van Wincoop (2003).

lateral imbalances, as it has been the case in the Eurozone, then the silver medal misspecification leads to an upward bias as the log of the sum (wrong procedure) overestimates the sum of the log (correct procedure). Finally, the bronze medal mistake concerns the price deflator: all the prices in the gravity equation are measured in terms of a common numeraire, so there is no price illusion. However, many authors deflate trade flows and GDP using the US CPI (following Rose's example). In this paper we include exports as dependent variable and define all the variables in nominal terms to avoid silver and bronze errors.

Finally, concerning the estimation problems, Santos-Silva and Tenreyro (2006) argue that the standard empirical methods used to estimate the gravity equation (i.e. Ordinary Least Squares, OLS) are inappropriate, even if these problems have been largely ignored by applied researchers, as the econometric methods commonly used to solve them were not easy to implement. Glick and Rose (2002) and Frankel and Rose (2002) exploited the time series information using panel data, giving birth to a literature in search of 'more reasonable' effects (Eicher and Henn, 2011). Micco *et al.* (2003) examined the dynamic impact of EMU on trade for 22 industrial countries using panel regressions based on a gravity model. Their findings suggest that EMU has fostered bilateral trade between 8% and 16% depending on the EMU membership of the countries and that the positive effect has been rising over time. Other studies, like Bun and Klaasen (2002) estimate a dynamic panel data model and distinguish between short (3.9%) and long-run effects (38%). Rose and Stanley (2005) perform a meta analysis of the results of 34 studies, and find a combined estimate of the trade effect between 30% and 90%, which is smaller than previous evidence. However, these papers generally use smaller and shorter datasets than Rose's. When they focus on large panels, they find bigger estimates (over 100%). Therefore, the empirical literature is far from conclusive and we can infer that dataset dimensions, and, especially, econometric approaches, influence the results.



While the heterogeneity bias is controlled through the use of fixed-effects, a second kind of misspecification is related to dynamics. The recent theoretical literature on international trade with heterogeneous firms (Bernard *et al.*, 2003; Melitz, 2003; Helpman *et al.*, 2004) has been largely based on evidence that, in a sector, the behaviour of firms can be highly heterogeneous, both concerning their productivity and their involvement in international transactions. In particular, the existence of sunk costs borne by exporters to set up distribution and service networks in the partner country may generate inertia in bilateral trade flows, especially among EMU countries, where there is also accumulation of invisible assets such as political, cultural and geographical factors characterizing the area and influencing the commercial transactions taking place within it.

Bun and Klaasen (2007) constitutes a path-breaking study in this respect. They show that the residuals of the Least Squares Dummy Variables estimator (LSDV) exhibit trends over time. Therefore, they estimate the gravity equation allowing for country pair specific time trends to account for the observed trending behaviour in the residuals. Moreover, they analyze the non-stationary nature of the data as well as the cointegration relationships and obtain a much smaller estimate of the euro effect (3%) on bilateral trade.<sup>4</sup> However, they employed methods that assume cross-section independence. The latter is an assumption unlikely to hold in bilateral trade data. As recently stated by Fidrmuc (2009), cross-correlation is likely to be present in gravity models because foreign trade is strongly influenced by the global economic shocks (i.e. other economies business cycles). Moreover, dependence is generated by construction as gravity models include bilateral trade flows together with aggregate national variables. Furthermore, the gravity model itself implies spatial dependence in the data due to the hypothesized effect of distance on trade. Several new panel unit root and cointegration tests have been proposed accounting for cross-sectional de-

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<sup>4</sup> Other papers that stress the importance of the non-stationary nature of the series and that apply cointegration techniques are Faruquee (2004) and Fidrmuc (2009).

pendence in the form of common factors.<sup>5</sup>

More recent studies have insisted on the importance of accounting for the existence of trends in the data and its possible non-stationary nature. Historically, researchers have assumed stationary time series to estimate gravity models. However, if the variables are non-stationary, a different statistical setup needs to be used. As Faruquee (2004) claimed, estimating the impact of a monetary union on trade faces several econometric challenges. Recent literature shows that the results of the gravity models are sensitive to their proper specification (Egger and Pfaffermayr, 2003). However, properly specified models in panel data may have some caveats when data are non-stationary. If the non-stationary nature of the series is not considered, spurious regressions may appear. Although the spurious correlation problem is less important in panels than in time series analysis, as the fixed effects estimator for non-stationary data is asymptotically normal (see Kao and Chiang, 2000), the results are biased. Correspondingly, panel cointegration techniques are used accounting for different possible estimation problems (endogeneity, cross-correlation or breaks). Therefore, a sound empirical strategy must proceed as follows: first, to determine the order of integration of the variables through panel unit root tests; second, to test for cointegration among the integrated variables using panel cointegration tests; finally, to use the panel cointegration estimators to provide reliable point estimates.

The contribution of our paper to the existing literature about the euro effect on trade is twofold. First, unlike previous research, (excepting Eicher and Henn, 2011) we address BT's critiques regarding the proper specification of gravity models and the definition of the variables, as we account for multilateral resistance, as well as unobserved bilateral heterogeneity. Second, we apply an econometric methodology comprising of a range of techniques to test and estimate efficiently in a non-stationary panel framework, solving endogeneity prob-

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<sup>5</sup> See for example Breitung and Pesaran (2008) for an overview of the literature and Gengenbach et al (2010) for a comparison of panel unit root tests.

lems as well as possible biases posed by structural breaks, country pair specific trends and cross-section dependence.

### **III. Data, methodology and empirical results**

Bun and Klaasen (2007) showed the importance of a correct specification of the gravity model including not only deterministic trend components but also stochastic trends derived from the non-stationary nature of the macro-variables involved. However, some practical problems implied that most of the evidence obtained so far did not considered nonstationarity. New developments in macroeconometrics have been recently extended to the panel framework allowing addressing most of the issues concerning both specification and estimation discussed in the previous section.

A first common problem in the context of panel non-stationary variables is that some widely used tests assume the absence of correlation across the cross-sections of the panel. That is, the individual members of the panel (countries) are considered independent. This assumption is not realistic and, therefore, cannot be maintained in the majority of the cases, especially when the countries are neighbours or are involved in integration processes. More recently, a second generation of panel tests, in contrast, introduce different forms of dependence, solving the above-mentioned problem.

Although there are several alternative proposals formulated in the literature to overcome the cross-section dependence problem, when dependence is pervasive –as in economic integrated areas- the best alternative is the use of factor models. This consists of assuming that the process is driven by a group of common factors, so that it is possible to distinguish between the idiosyncratic component and the common component.

In the case of panel unit roots, several tests have been formulated based on factor mod-

els.<sup>6</sup> In particular, Bai and Ng (2004) account for the non-stationarity of the series coming either from the common factors, the idiosyncratic component or from both. Moreover, they consider the possible existence of multiple common factors as well as the existence of cointegration relationships among the series of the panel. Banerjee *et al.* (2004) stated that there is a tendency to over-reject the null of stationarity when cointegration is present. As the existence of cointegrating relations between trade series is a very plausible hypothesis in economic integrated areas, the proposal in Bai and Ng (2004) is the best approach in our case.<sup>7</sup> For the sake of comparison, we will also present the results obtained using Pesaran's (2007) approach. Similarly, we will also allow for dependence in the estimation of the cointegration relationships using the common factor approach of Bai and Ng (2004).

A second caveat appears when there are structural breaks in the time dimension of the panel. If there exist linear combinations of integrated variables that cancel out their common stochastic trends then, these series are said to be cointegrated. The economic translation is that these series share an equilibrium relationship. However, a commonly neglected phenomenon is that both, the cointegrating vector and the deterministic components might change during the period analyzed, and if we do not take account of these structural breaks in the parameters of the model, inference concerning the presence of cointegration can be affected by misspecification errors. Therefore, in this paper we propose the use of the tests developed in Banerjee and Carrion-i-Silvestre (2010). They generalize the approach in Pedroni (1999, 2004) to account for one structural break that may affect the long run relationship in a number of different ways (cointegrating vector and/or deterministic components). Moreover, they address the cross-section dependence issue by using the above-mentioned factor model approach due to Bai and Ng (2004) to generalize the degree of per-

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<sup>6</sup> Namely, Pesaran (2007), Phillips and Sul (2003), Moon and Perron (2004) and Bai and Ng (2004).

<sup>7</sup> Moreover, using Monte Carlo methods, Gengenbach *et al.* (2010) and Jang and Shin (2005) show that, for all the specifications considered in their simulation experiments, the test in Bai and Ng (2006) has more power than those by Moon and Perron (2004) and Pesaran (2007), and better empirical size than that of Phillips and Sul (2003).

missible cross-section dependence allowing for idiosyncratic responses to multiple common factors.

To sum up, we control for econometric issues usually neglected in earlier literature: first, we account for cross-section dependence among countries in the panel tests, both unit roots and cointegration. Second, we allow for the existence of a break in the deterministic components (constant, trend and cointegrating vector) of the model as well as in the cointegration relationship, a major point to assess the effect of institutional changes in the relationship. Furthermore, since the trend included in the specification is country pair specific, the break in the trend is also allowed to have different coefficients for each country pair, therefore allowing for a higher degree of heterogeneity in the estimation. To the best of our knowledge, this is the first time that structural changes have been considered in the euro effect literature based on gravity equations. Finally, the estimation of the long-run relationship uses a methodology that not only efficiently estimates the coefficients but also is based on the common factors decomposition that assures a homogeneous econometric approach. We choose, for this purpose, the CUP Fully Modified (CUP-FM) and CUP Bias Corrected (CUP-BC) estimators by Bai *et al.* (2009).

## **Data**

The countries included in the study are Australia, Austria, Belgium and Luxembourg (as an unique area), Canada, Chile, Denmark, Finland, France, Germany, Greece, Iceland, Ireland, Italy, Japan, South Korea, Mexico, Netherlands, New Zealand, Norway, Poland, Portugal, Spain, Sweden, Switzerland, United Kingdom and United States.

The dataset contains annual data from these 26 OECD countries and covers the period 1967-2008. Although the number of years available was higher, we have opted by restrict our sample to this period, in order to exclude the effects of the financial crisis that started in

2008. Hence, we have a balanced panel with dimension  $N=650$  (all possible bilateral combinations of countries) and  $T=42$ . The total number of observations is  $NT=27,300$ .

Following the discussion in section 2, one of the contributions of the paper is to perform the analysis and the estimation of the gravity equation for the euro effect using the variables correctly defined. Concerning the dependent variable, we include exports as dependent variable instead of the average of exports and imports, as it is frequently made in the literature. As BT points out, the gravity equation is an expenditure function that explains uni-directional bilateral trade flows. De Benedictis and Taglioni (2011) also reinforce this point, arguing that the choice of the dependent variable should be driven by theoretical considerations, which privilege the use of uni-directional trade data.<sup>8</sup> Hence,  $EXPORTS_{ijt}$  is to the log of the export flows from country  $i$  to country  $j$  in nominal terms<sup>9</sup> - instead of real terms, according to BT's critiques- and  $GDP_{it}$  and  $GDP_{jt}$  are the nominal GDPs in the logs of the exporter and importer country respectively, obtained from the CHELEM – CEPII database and expressed in current dollars. Additionally, two dummy variables have been built to include the effect of particular integration agreements on trade. Namely  $RTA_{ijt}$  which is 1 if both countries have a free trade agreement at time  $t$  and is constructed using World Trade Organization (WTO) data, and finally the key variable of interest  $EURO_{ijt}$  which equals 1 if both trading partners belong to the euro area in year  $t$  and zero otherwise. To the extent that these agreements are made or dissolved during the sample period, this variable is distinct from the time-invariant country-pair fixed effect.

The formal model that we estimate comes from the gravity equation, and in particular, we follow the traditional specification from the recent literature on the euro effect using non-stationary panels. The purpose is to isolate the effects of EMU on exports trying to control

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<sup>8</sup> See De Benedictis and Taglioni (2011), p. 71.

<sup>9</sup> Since we include OECD countries, the total number of zero observations represents only the 0.2% of total flows (64 observations). We have replaced these zero flows by 0.01.

for other factors that may have an influence on exports flows but are not related to the monetary union. The gravity model predicts that bilateral exports should depend on factors such as economic size or ‘mass’, distance, and other related considerations. Bearing this in mind the basic panel equation in the literature can be expressed as follows:

$$EXPORTS_{ijt} = \beta_1 GDP_{it} + \beta_2 GDP_{jt} + \delta_1 RTA_{ijt} + \delta_2 EURO_{ijt} + \eta_{ij} + \tau_{ij} \cdot t + \varepsilon_{ijt} \quad (1)$$

where  $\eta_{ij}$  is a country specific fixed effect,  $\tau_{ij} \cdot t$  is a country pair specific time trend and  $\varepsilon_{ijt}$  is the error term.<sup>10</sup>

The fixed effect ( $\eta_{ij}$ ) is intended to capture all individual fixed factors, including unobservable characteristics associated with a given pair of countries that have affected bilateral trade flows historically. These time invariant factors include geographical distance, area, common language, common border, etc. The advantage of fixed effects estimation over directly including these specific measures is controlling for omitted variables bias as a whole at the expense of isolating the individual contribution of each of the variables considered (Micco et al, 2003).<sup>11</sup>

The country pair specific time trend,  $\tau_{ij} \cdot t$ , is intended to capture all country-pair specific omitted trending variables, for instance, institutional characteristics, factor endowments, and cultural aspects that may change over time.<sup>12</sup> Therefore, the approach that we follow to account for trend effects is very flexible and considers both, the time dimension and the heterogeneous behaviour (coefficients) across country-pairs. Potential bias due to the existence

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<sup>10</sup> Later in the analysis, we will include additional deterministic trends in equation (1), which correspond to structural breaks in the constant, the trend or both.

<sup>11</sup> Hence, the model does not include distance between countries as an explanatory variable and assume that country-pair specific fixed effects will account for the distance effect. Moreover, as we have previously stated, the econometric approach used in this paper accounts for spatial dependence properly.

<sup>12</sup> Country-pair specific variables, such as transport costs or tariff, can vary over time due to technical progress in transport and telecommunications or to the trade liberalization process, generating trends in trade that must be accounted for.

of common time effects is also controlled through the use of common factors; hence, time effects are not included in the specification.

The set of coefficients  $\delta_l$  and  $\delta_2$  represents the effect of any free trade agreements and EMU on member states' exports to their country peers (including extra-area trade). Therefore, the parameter of interest is  $\delta_2$  and the difference in exports before and after the introduction of the euro is used to identify this coefficient.

The next subsections are devoted to the presentation of the empirical results, comprising panel estimates of the EMU trade effects at the area-wide level as well as cross-country differences.

### **Panel unit root, stationarity tests and cross-section dependence**

We use a testing procedure that deals with the problem of cross-section dependence. First, we compute the test statistic by Pesaran (2004) to assess whether the time series in the panel are cross-section independent. Then, we proceed in a second stage to compute unit root statistics that account for such dependence when required.

Pesaran (2004) proposes a test statistic based on the average of pair-wise Pearson's correlation coefficients of the residuals obtained from an autoregressive (AR) model. Under the null hypothesis of cross-section independence the statistic converges to a standard normal distribution. The results in Table 1 show that the null hypothesis of independence is strongly rejected in the case of exports, so that cross-section dependence has to be considered when computing the panel data statistics if misleading conclusions are to be avoided. Note that, according to Pesaran (2004) the CD test is valid for  $N$  and  $T$  tending to  $\infty$  in any order and that it is particularly useful for panels with small  $T$  and large  $N$ . Moreover, this test is also robust to possible structural breaks, which makes it especially suitable for our study.



TABLE 1

*Pesaran's CD and CADF statistics*

<i>Variable</i>	<i>CD dependence test</i>	<i>CADF panel unit root test</i>
GDP <sub>it</sub>	-0.146	-1.235
GDP <sub>jt</sub>	-0.146	-1.235
EXPORTS <sub>ijt</sub>	105.136***	-0.964

*Notes:* \*\*\* denotes rejection at 1% level. All variables are in logarithms. One lag is selected for real and nominal GDP; two lags for nominal exports according to BIC criterion. Trend and constant are included in all cases.

Once we have found evidence of dependence, we study the order of integration of the variables. We follow Pesaran (2007) and Bai and Ng (2004) and specify the unit root tests allowing for cross-sectional dependence as driven by a common factor model, so that it is possible to distinguish between the idiosyncratic component and the common component. While Pesaran (2007) focuses on the extraction of the common factors that generate the cross correlations in the panel to assess the non-stationarity of the series, in Bai and Ng (2004) the non-stationarity of the series can come either from the common factors, the idiosyncratic component or from both. Moreover, Pesaran (2007) only considers the existence of one common factor<sup>13</sup> while the other alternative can consider several ones. We implement both tests in this section. The results obtained from the Pesaran CADF test are reported in Table 1 concluding in favour of non-stationarity, with a critical value of -2.50 at a 5% confidence level.

Bai and Ng (2004) approach allows to control for cross-section dependence given by cross-cointegration relationships, potentially possible among our group of countries and variables. For the estimated idiosyncratic component, they propose an ADF test for individual unit roots and a Fisher-type test for the pooled unit root hypothesis ( $P_e$ ), which has a

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<sup>13</sup> The main advantage of this method is its simplicity to compute while its drawback is that the behaviour of the idiosyncratic component is to some extent neglected being assumed its stationarity.

standard normal distribution. The estimation of the number of common factors is obtained using the panel BIC information criterion as suggested by Bai and Ng (2002), with a maximum of six common factors. Results are summarized in Table 2. Concerning the idiosyncratic component, the results of the panel ADF unit root tests clearly point to the rejection of the unit root hypothesis; however, the results of the unit root analysis of the factor component for all the variables analyzed point to nonstationarity. In none of the cases presented in Table 2 can the null hypothesis of independent stochastic trends be rejected. Thus, the variables are nonstationary and its source is not variable-specific, but associated to the common factors.

TABLE 2

*Panel Data Statistics based on Approximate Common Factor Models  
Bai and Ng (2004) statistics*

	$EXPORTS_{ijt}$	$GDP_{it}$	$GDP_{jt}$
Idiosyncratic ADF statistic	-3.387***	-1.849***	-1.849***
MQ test (parametric)	-34.968	-21.987	-21.987
MQ test (non-parametric)	-32.057	-23.343	-23.343

*Notes:* \*\*\* denotes rejection at 1% level. The tests on the factors are asymptotically independent of the tests on the idiosyncratic errors.  $MQ_c$  and  $MQ_f$  use a non-parametric and a parametric correction respectively to account for additional serial correlation. Both statistics have a non-standard limiting distribution.

### Panel cointegration

As in the case of the unit root tests, the main caveat of the first generation panel cointegration tests is that they do not consider the presence of cross-section dependence among the members of the panel.<sup>14</sup> Trying to solve the problem of cross-section dependence, new statistics have been also designed to test for cointegration, using factor models in a fashion similar to the one proposed by Bai and Ng (2004) for unit root testing. Moreover, as

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<sup>14</sup> We have also applied the panel cointegration tests proposed by Kao (1999) and McCoskey and Kao (1998) for the sake of comparison. The complete results are available from the authors upon request.

the existence of structural breaks in the cointegrating relationships biases the results in panel settings –see Banerjee and Carrión-i-Silvestre (2010) – they propose an extension of the Gregory and Hansen (1996) approach using common factors to account for dependence. Banerjee and Carrion-i-Silvestre (2010) propose a panel test for the null hypothesis of no cointegration allowing for breaks both in the deterministic components and in the cointegrating vector and also accounts for the presence of cross-section dependence using factor models.

In Table 3 we present the results of the test. We apply the statistics based on the accumulated idiosyncratic components,  $Z_j^*$  for the eight potential specifications allowed by the test.<sup>15</sup> With all of them the null hypothesis of non-cointegration is rejected. Using the BIC information criterion we choose specification 5, which contains a constant and a trend and a structural break that affects them both simultaneously. The date of the break is found in 1989.

TABLE 3

*Banerjee and Carrion (2010) BC cointegration tests*

<i>Model</i>	$Z^*$	$r$	$r_1$
1	-5.66***	6	6
2	-5.59***	6	6
3	-7.72***	6	6
4	-6.19***	6	6
<b>5</b>	<b>-15.88***</b>	<b>6</b>	<b>6</b>
6	-10.02***	6	6
7	-16.09	6	6
8	-15.97	6	6

*Notes:* \*\*\* denotes rejection at 1% level. Specification 5 is selected according to the BIC criterion; it includes a constant, a trend and a break in both components. The break takes place in 1989. The null of no cointegration is rejected in all cases.  $r_1$  is the number of independent stochastic trends underlying the  $r$  common factors;  $r$  is the total number of factors allowed in the specification.

<sup>15</sup> See the appendix for further information about the test.

### Estimation of the gravity equation

Once the different tests applied have provided us with evidence of cointegration, either considering a stable relationship or instabilities, we should obtain the long-run estimates using consistent techniques. Kao and Chiang (2000) recommended the fully modified (FM) estimator of Phillips and Hansen (1990) and the dynamic ordinary least squares (DOLS) estimator as proposed by Saikkonen (1991) and Stock and Watson (1993). However, although both consistently estimate the long-run parameters and correct for autocorrelation and endogeneity, they do not account for dependence. Alternatively, Bai *et al.* (2009) consider the problem of estimating the cointegrating vector in a cointegrated panel data model with non-stationary common factors. They treat the common  $I(1)$  variables as parameters. These are estimated jointly with the common slope coefficients  $\beta$  using an iterated procedure. Although this procedure yields a consistent estimator of  $\beta$ , the estimator is asymptotically biased. To account for this bias, the authors construct two estimators that deal with endogeneity and serial correlation and re-center the limiting distribution around zero. The first one, CUP-BC, estimates the asymptotic bias directly. The second, denoted CUP-FM, modifies the data so that the limiting distribution does not depend on nuisance parameters. Both are ‘continuously-updated’ (CUP) procedures and require iteration till convergence. The estimators are  $\sqrt{nT}$  consistent and enable the use of standard tests for inference. Finally, the approach is robust to mixed  $I(1)/I(0)$  factors as well as mixed  $I(1)/I(0)$  regressors.

We present in Table 4 the results of the CUP estimation using the methodology of Bai *et al.* (2009). We have based our estimation on the results previously obtained using the Banerjee and Carrión-i-Silvestre (2010) tests concerning not only the cointegration tests, but also the deterministic specification of the chosen model. Bai *et al.* (2009) consider extensions of their estimators when the assumptions about the deterministic components are relaxed. In order to account for the existence of incidental trends (intercept and/or trend), they define

accordingly a projection matrix  $M$  for demeaned and/or detrended variables. We concentrate the deterministic components before we estimate the long-run parameters. As we have mentioned before, among the deterministic components we include the constant, the country pair specific trends, the common break in the constant and the common break in the country pair specific trends.<sup>16</sup> The number of common factors for the estimation is selected according to Principal Components Factor Analysis (PCA henceforth).

Therefore, once we have performed this transformation we are able to apply the Bai *et al.* (2009) estimators. The results are shown in Table 4, where we have also included the LSDV estimation results and the Bai and Ng (2006) two-step fully-modified estimator (Bai FM henceforth) for the sake of comparison. However, it should be noted that the only estimators that are consistent when the common factors are non-stationary are the CUP-FM and the CUP-BC. These results are presented in the last two columns of the table. Although the LSDV estimator is the most commonly applied in the gravity literature, the parameters obtained are biased when the common factors are non-stationary. The size of this bias is shown in Bai *et al.* (2009) and this may explain earlier results in the applied literature.

As mentioned above, the variables are constructed according to BT's critiques. We have transformed them to account for the deterministic components and the structural break found in 1989. The EMU dummy is correctly signed and significant. The CUP-BC and CUP-FM estimators provide lower results than LSDV and BaiFM, which confirm our theoretical predictions of the need of accounting for dependence and nonstationarities. We should note that LSDV estimator is shifted away from zero due to the asymptotic bias induced by the cross-section dependence. The RTA coefficient is positive and significant and its effect is also notably reduced when using the proper estimators.

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<sup>16</sup> Note that this implies that in the model specification of the gravity equation in expression (1) above, we have filtered the five variables (EXPORTS, GDPi, GDPj, RTA and EMU) of the deterministic components.

TABLE 4

*CUP estimation of the long-run parameters 1967-2008*

<i>Variables</i>	<i>LSDV</i>	<i>Bai FM</i>	<i>CUP-FM</i>	<i>CUP-BC</i>
$GDP_{it}$	1.17*** (64.00)	0.67*** (27.14)	0.64*** (25.54)	0.64*** (25.37)
$GDP_{jt}$	1.08*** (59.66)	0.79*** (27.18)	0.78*** (26.34)	0.78*** (26.29)
RTA	0.79*** (13.41)	0.33*** (7.55)	0.22*** (5.22)	0.22*** (3.36)
EMU	0.56*** (4.23)	0.26*** (3.36)	0.17** (2.23)	0.16** (2.07)

*Notes:* Bold letters indicate significance at a 5% level. The specification 5 is estimated with 2 common factors according to PCA. Results with a different number of factors are available under request. The common structural break takes place in 1989. The t-statistic is reported in parenthesis.

Concerning the GDP variables, the values obtained are around 0.65 and 0.8 for the exporter and importer respectively. The importer GDP shows a higher coefficient than the exporter GDP, indicating that demand has a greater influence on exports than supply. Again, the two estimated coefficients obtained using LSDV are much larger than those obtained with the other estimators due to the above-mentioned bias. The Bai FM estimator, in contrast, corrects for the presence of dependence and assumes stationary common factors. However, Bai *et al.* (2009) strongly recommend the use of the CUP-FM and CUP-BC when there is dependence and the common factors are non-stationary. The common structural break occurs in 1989. We attribute this break to the effects of the Single European Act, which was signed in 1987.

Therefore the main empirical findings can be summarized as follows: first, there exists a long-run relationship linking trade and the gravity equation variables in a system that exhibits cross-section dependence and non-stationary common factors, which cancel-out in cointegration. Second, there are some significant instabilities (structural breaks) that can be identified using panel cointegration tests that also account for the common factors. Third, the ex-

istence of dependence and non-stationary common factors makes it necessary to use consistent estimators, notably the CUP-FM and CUP-BC estimators proposed by Bai *et al.* (2009). All in all, the unrealistically high effects of the euro on trade found in previous empirical literature mostly disappear when the trend of the integration process is accounted for. Our results are in line with the most recent literature started with Bun and Klaasen (2007), Fidmurt (2009), Gengenbach (2009) and Berger and Nitsch (2008). They show that the increase in trade within the euro area is simply a continuation of a long-run trend, probably linked to the broader set of EU's economic integration policies and institutional changes, the euro having just a residual effect.

#### **4. Summary and concluding remarks**

In this paper we try to fill the gaps present in the previous literature on euro effects on trade. Using a data set that includes 26 OECD countries from 1967 to 2008, we estimate a gravity equation through a cointegration approach fully allowing for cross-section dependence. The analysis consists of three steps. First, unit root tests for cross-sectionally dependent panels are applied. Second, the existence of a cointegration relationship among the variables of a proper specification of the gravity equation is tested. In this exercise we account both for dependence in the cross-section dimension and discontinuities in the deterministic and the cointegrating vector in the time dimension. Third, the appropriate CUP-BC and CUP-FM estimators are used to estimate the long-run relationships.

To the best of our knowledge, this is the first attempt to jointly incorporate in the estimation of a gravity equation for the assessment of the euro effect the following aspects: first, we include Baldwin's critiques in terms of model specification and variables' construction and we include country-pair specific trends; second, we account for the existence of cross-sections dependence as well as structural breaks in the time domain; and third, we consider the non-stationary nature of the series involved in the analysis. This approach allows us to

put the adoption of the euro by EMU members in historical perspective. We argue that the creation of the EMU is best interpreted as a progression of policy changes that have contributed to greater economic integration among EMU countries over the last decades. We find strong evidence of a gradual increase in trade intensity between European countries as well as pervasive cross-section dependence. Once we control for both, dependence and this (breaking) trend in trade integration, the effect of the formation of the EMU mostly fades out in line with most recent empirical literature.



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### Appendix A: Banerjee and Carrion-i-Silvestre (2010) test

Banerjee and Carrion-i-Silvestre (2010) propose a panel test for the null hypothesis of no cointegration allowing for breaks both in the deterministic components and in the cointegrating vector and also accounts for the presence of cross-section dependence using factor models. They define a  $(m \times I)$  vector of non-stationary stochastic process,  $Y_{i,t} = (y_{i,t}, x'_{i,t})$  whose elements are individually I(1) with the following Data Generating Process:

$$y_{i,t} = D_{i,t} + x'_{i,t} \delta_{i,t} + u_{i,t} \quad (2)$$

The general functional form for the deterministic term  $D_{i,t}$  is given by:

$$D_{i,t} = \mu_i + \beta_i t + \sum_{j=1}^{m_i} \theta_{i,j} DU_{i,j,t} + \sum_{j=1}^{m_i} \gamma_{i,j} DT_{i,j,t}, \quad (3)$$

where  $DU_{i,j,t} = 1$  and  $DT_{i,j,t} = (t - T_{i,t}^b)$  for  $t > T_{i,t}^b$  and 0 otherwise,  $T_{i,t}^b = \lambda_{i,j}^b T$  denotes the timing of the  $j$ -th break,  $j = 1, \dots, m_i$ , for the  $i$ -th unit,  $I = 1, \dots, N$ ,  $\lambda_{i,j}^b T \in \Lambda$ , being  $\Lambda$  a closed subset of  $(0,1)$ . The cointegrating vector is a function of time so that

$$\delta_{i,t} = \begin{cases} \delta_{i,1} T_{i,0}^c < t \leq T_{i,1}^c \\ \delta_{i,2} T_{i,1}^c < t \leq T_{i,2}^c \\ \dots\dots \\ \delta_{i,j} T_{i,j-1}^c < t \leq T_{i,j}^c \\ \dots\dots \\ \delta_{i,n_i+1} T_{i,n_i}^c < t \leq T_{i,n_i+1}^c \end{cases} \quad (4)$$

with  $T_{i,0}^c = 0$  and  $T_{i,n_i+1}^c = T$ , where  $T_{i,j}^c = \lambda_{i,j}^c T$  denoting the  $j$ -th time of the break,  $j = 1, \dots, n_i$ , for the  $i$ -th unit,  $i = 1, \dots, N$ , for the  $i$ -th unit,  $i = 1, \dots, N$ ,  $\lambda_{i,j}^c \in \Lambda$ .

Banerjee and Carrion-i-Silvestre (2010) propose eight different model specifications:

**Model 1.** Constant term, no linear trend -  $\theta_{ij} = \beta_i = \gamma_{ij} = 0 \quad \forall i, j$  in (3) – and constant cointegrating vector.

**Model 2.** Stable trend -  $\theta_{ij} = 0; \beta_i \neq 0 \quad \forall i$  and  $\gamma_{ij} = 0 \quad \forall i, j$  in (3) – and constant cointegrating vector.

**Model 3.** Constant term with shifts; stable trend -  $\theta_{ij} \neq 0; \beta_i \neq 0; \gamma_{ij} = 0 \quad \forall i, j$  (3) – and constant cointegrating vector. The model considers multiple level shifts.

**Model 4.** Constant term, trend and changes in trend, -  $\theta_{ij} = 0; \beta_i \neq \gamma_{ij} \neq 0 \quad \forall i, j$  in (3) – and constant cointegrating vector. The model considers multiple trend shifts.

**Model 5.** Changes in constant and trend -  $\theta_{ij} \neq 0; \beta_i \neq 0$  and  $\gamma_{ij} \neq 0 \quad \forall i, j$  in (3) – and constant cointegrating vector. The model considers multiple trend and level shifts.

**Model 6.** No trend, constant term with shifts -  $\theta_{ij} \neq 0; \beta_i = 0 \quad \forall i, j$  in (3) – and changes in the cointegrating vector.

**Model 7.** Constant term, trend; changes in the level -  $\theta_{ij} \neq 0; \beta_i \neq 0 \quad \forall i, j$  in (3) – and changes in the cointegrating vector.

**Model 8.** Constant term, trend; changes in the level and the trend -  $\theta_{ij} \neq 0; \beta_i \neq 0$  and  $\gamma_{ij} \neq 0 \quad \forall i, j$  in (3) – and changes in the cointegrating vector

The common factors are estimated following the method proposed by Bai and Ng (2004). They first compute the first difference of the model; then, they take the orthogonal projections and estimate the common factors and the factor loadings using principal components.

In any of these specifications, Banerjee and Carrion-i-Silvestre (2010) recover the idiosyncratic disturbance terms ( $\tilde{e}_{i,t}$ ) through cumulation of the estimated residuals and propose testing for the null of no cointegration against the alternative of cointegration with break us-

ing the ADF statistic.

The null hypothesis of a unit root can be tested using the pseudo  $t$ -ratio  $t_{\tilde{\epsilon}_i}^j(\lambda_i)$ ,  $j = c, \tau, \gamma$ .

The models that do not include a time trend (Models 1 and 6) are denoted by  $c$ . Those that include a linear time trend with stable trend (Models 2, 3 and 7) are denoted by  $\tau$  and, finally,  $\gamma$  refers to the models with a time trend with changing trend (Models 4, 5 and 8). When common (homogeneous) structural breaks are imposed to all the units of the panel (although with different magnitudes), we can compute the statistic for the break dates, where the break dates are the same for each unit, using the idiosyncratic disturbance terms.